NBER WORKING PAPER SERIES

DEMOGRAPHIC CHANGE AND THE STRUCTURE OF WAGES: A DEMAND-THEORETIC ANALYSIS FOR BRAZIL

Ernesto F. L. Amaral
Daniel S. Hamermesh
Joseph E. Potter
Eduardo L.G. Rios-Neto

Working Paper 13533
http://www.nber.org/papers/w13533

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
October 2007

We gratefully acknowledge financial support from the John D. and Catherine T. MacArthur Foundation, the William and Flora Hewlett Foundation, and the Population Reference Bureau. Viviana Salinas and Eunice Vargas played an important role in the discussions that led up to this analysis, Starling Pullum helped set up our databases, and helpful comments were received from David Weil and participants at several conferences and universities. The views expressed herein are those of the author(s) and do not necessarily reflect the views of the National Bureau of Economic Research.

Demographic Change and the Structure of Wages: A Demand-Theoretic Analysis for Brazil
Ernesto F. L. Amaral, Daniel S. Hamermesh, Joseph E. Potter, and Eduardo L.G. Rios-Neto
NBER Working Paper No. 13533
October 2007
JEL No. J23,O15

ABSTRACT

With rapidly declining fertility and increased longevity the age structure of the labor force in developing countries has changed rapidly. Changing relative supply of workers by age group, and by educational attainment, can have profound effects on labor costs. Their impacts on earnings have been heavily studied in the United States but have received little attention in Asia and Latin America, where supply shocks are at least as large and have often proceeded less evenly across the economy. We use data on 502 local Brazilian labor markets from Censuses 1970-2000 to examine the extent of substitution among demographic groups as relative supply has changed. The results suggest that age-education groups are imperfect substitutes, so that larger age-education cohorts see depressed wage rates, particularly among more-educated groups. The extent of substitution has increased over time, so that the decreasing size of the least-skilled labor force today is barely raising its remaining members' wages.

Ernesto F. L. Amaral
Department of Demography (CEDEPLAR)
Federal University of Minas Gerais
Belo Horizonte, Minas Gerais BRAZIL
amaral@cedeplar.ufmg.br

Daniel S. Hamermesh
Department of Economics
University of Texas
Austin, TX 78712-1173
and NBER
hamermes@eco.utexas.edu

Joseph E. Potter
Department of Sociology
University of Texas
Austin, TX 78712
joe@prc.utexas.edu

Eduardo L.G. Rios-Neto
Department of Demography (CEDEPLAR)
Federal University of Minas Gerais
Belo Horizonte, Minas Gerais BRAZIL
eduardo@cedeplar.ufmg.br

I. Introduction

The substitution of one group of workers for another as their relative supplies change has received substantial attention in the literature on labor demand (Hamermesh 1993, Chapter 3). This attention is justified by the importance of the potential effects on wage rates and earnings and their role in such areas on which economic and social policy might focus as old-age assistance, payroll taxes and labor productivity. The topic received attention in wealthy countries generally (Bloom *et al* 1987) and in the United States in particular in the 1970s and 1980s as a result of concerns about youth wages after the baby boom of 1946-62 (Freeman 1979; Welch 1979; Berger 1985; Triest *et al* 2006).

While the impact of demographic change on the structure of wages was shown to be at least temporarily substantial in the U.S.; and while it has no doubt mattered (probably to a lesser extent) in other rich countries since 1945, wealthy countries may not be the best laboratories for examining it. In the U.S. the share of the male population ages 15-24 increased from 13.5 percent in 1960 to 19.5 percent in 1980. In Brazil, for example, the rise in longevity and decline in fertility caused a similar percentage-point drop over two decades, from 37.0 percent in 1980 to 31.7 percent in 2000. Changes in the skill structure of the labor force may be even larger and more rapid: The largest decadal change in the percentage of American men with at least a college education was 6.6 percentage points between 1970 and 1980. In Brazil between 1991 and 2000 the percentage of adult men with nine or more years of schooling rose from 20.6 to 29.7. At least as important, the timing of changes in both the age and educational structures of the labor force in developing countries, including Brazil, has been much more geographically heterogeneous than in richer countries. These sorts of changes in the structure of the labor force in developing countries may thus provide greater scope for identifying the impacts of changes in relative supply on earnings than the demographic shocks of the past half century in richer economies.

Despite the importance of the topic for economic development and the possibly greater variability in the forcing variables in developing countries that would facilitate identification of earnings effects, no studies of developing countries have examined the role of these supply shocks in a formal model of labor demand. Careful research on the role of changing education and age endowments on wage inequality has been conducted (Gindling and Robbins 2001), and other studies have examined labor demand by firms in Latin America (e.g., Roberts and Skoufias 1997; Saavedra and Torero 2004); but no study has considered the impact of demographic change in a complete model of labor demand. We contribute to the literatures on demographic change in developing countries, and to the general study of labor demand, by including cohort size and the shifting structure of educational attainment in a standard model. Implicitly we assume throughout that relative supplies of workers classified by age/education are exogenous to the firm and that firms bid for workers based on their relative scarcity. Using the example of Brazil, the result is a set of estimates of factor-price elasticities by age-education group in an economy in which relative supplies of workers are changing rapidly and differentially across labor markets.

II. Demographic Dividends and Labor Demand

The topic we address is related to, but different in its focus from recent studies of the "demographic dividend"—how changing age structures in developing countries resulting from sustained and rapid fertility decline present a temporary "window of opportunity" during which the reduced dependency ratio can yield high rates of growth in per-capita income. This positive impact derives from the mechanical link between the sizes of the working age and total populations, increases in labor supply due to higher proportions of women becoming employed, higher savings rates, higher rates of human capital formation, and, possibly, from the effect of population aging on capital accumulation via capital deepening. That the decline in the dependency ratio caused by rapid fertility decline has substantially influenced economic development in East and Southeast Asia has been shown often (e.g., Bloom and Freeman 1986; Bloom *et al* 2003; Williamson 2003; Mason 2005). These authors stress the transitory nature of the decrease in the dependency ratio, and the conditional impact of the dividend: The drop in the dependency ratio will only result in economic growth in the right policy environment.

While the demographic dividend literature focuses on the ratio of the population of working age to the rest of the population, the analysis here focuses on the changes in age structure *within* the working-age population that necessarily accompany the demographic transition, as well as the concurrent changes in education levels that normally accompany and may even drive the demographic transition. While in the dividend literature the focus is mainly on aggregate outcomes, our concern is with the distribution of economic outcomes within the labor force and how these may be affected by its changing composition.

Evidence from studies of the U.S. baby boom suggests that increases in factor supplies led to declines in wage rates of the expanding sub-aggregates, confirming the role of negative own-quantity elasticities of factor price. Similarly, one would expect that an increase in the supply of skilled labor will lead to a relative decline in its wage rate. In developed countries, however, this decline has not been observed, with the skill premium actually increasing along with the supply of educated workers (Katz and Murphy 1992; Autor *et al* 1998). The reasons may be skill-biased technical change, the role of international trade, or others; but the results require expanding the usual production-function framework.

Despite their difficulties in describing aggregate trends over the past three decades, studies of labor-labor substitution in rich countries illustrate the power of a formal factor-demand framework and the richness of combining age (experience) and schooling as basic labor inputs driving variations in wage rates. While the technology-constant microeconomic findings are swamped at the macro level by trends arising from aggregate shocks, they support the basic tenets of production theory. Using data at the micro level from a developing country, we should be able to take a formal model and estimate more precisely how changing cohort size and skill alter relative wage differentials.

III. Brazilian Background and Data

In Brazil the shocks that generated subsequent changes in relative labor supply by cohort began in the early 1960s with a decline in fertility in the metropolitan areas of Rio de Janeiro, São Paulo, and Porto Alegre, which had total fertility rates below five. From there the decline spread to the interior of these Southeastern states and to the capital cities of states in the Central-West, North and Northeast, finally reaching the interior and rural areas of those regions in the 1980s. In 2000 a substantial number of municipalities still had total fertility rates above four, while there were also many where fertility had fallen below replacement. The differences in the timing and speed of the fertility transition led to substantial differences in age distributions across states and municipalities and also to differential changes in these distributions at different points in time (Potter *et al* 2002).

The longest series on age, education and earnings come from the Brazilian Censuses conducted in 1960, 1970, 1980, 1991, and 2000. Microdata from these Censuses are available from long-form questionnaires administered to 25-percent samples in 1960, 1970 and 1980. In 1991 and 2000 the sample sizes depended on the size of the municipality, with 10-percent samples from municipalities having more than 15,000 inhabitants, and 20-percent samples from smaller municipalities. In all cases there are records for every individual in the sampled households, containing information on age, gender, marital status, educational attainment, school enrollment, and, if employed, occupation and earnings. There are also questions on migration, including state of birth, previous residence, and residence five years before the Census.

The lowest level of geographic identifier on these records common to all Censuses is the *município*, since information on *distritos*, the sub-divisions of *municípios*, is not available in the Census microdata. In previous work, Potter *et al* (2002) established minimum comparable areas that account for the changing definitions and divisions of *municipios* across the Census years, which is necessary since their number increased from approximately 2,300 in 1960 to 5,280 in 2000. They were able to aggregate minimum comparable areas into 502 micro-regions across the five Censuses.[1] It is thus possible to calculate various statistics summarizing the age distributions, labor-market outcomes and education indicators for each of these 502 consistently-defined areas in each of the five Censuses. In the end, because the 1960 Census categorized earnings by bracket, we exclude it and base the analyses on the Censuses beginning 1970. Using these very small geographical units, however, poses the question of internal migration, which has not been incorporated in most previous analyses undertaken at the national level.[2] We discuss this potential problem in more detail later.

We categorize the labor force into four age groups: Youths (15-24), young adults (25-34), experienced adults (35-49) and older adults (50-64). The widths of the age categories are unequal in order to make the fractions in each category somewhat more equal. In light of the massive evidence on the role of education in defining sub-aggregates of labor (Hamermesh 1993, Chapter 3; Borjas 2003), we cross-classified workers by their educational attainment based on Riani (2005). She notes that by 2000 the majority of Brazilians between ages 7 and 14 were in school, and large fractions were completing elementary school. Moreover, she indicates that there was a decrease in regional, racial and rural-urban differentials in elementary school attainment. On the other hand, although more people are attending secondary school, the proportion of people having between nine and twelve years of schooling is still small, and regional differences are still substantial. Taking into account the specifics of the Brazilian population, we thus classify workers by educational attainment into three main groups: Zero to four years, five to eight years, and at least nine years of schooling.

Because of differences in educational attainment by cohort, using separate vectors of indicators for age and education misses the rapid change in educational attainment across cohorts

---

[1]Note that these micro-regions differ from those defined by the Brazilian Institute of Geography and Statistics (IBGE) and available in the Census microdata, but closely approximate those defined for the 1991 Census.

[2]See Borjas *et al* (1997) for a discussion of the role of internal migration in modifying the impacts of exogenous changes in relative supply on relative wage rates.

and would generate multicollinearity when estimating factor-price elasticities. To obviate this difficulty we generate a full set of interactions of the four age indicators and the three education categories. For each of the twelve cells, for each micro-region and each year we calculated mean earnings and the proportion of males in each age-education group. The dependent variable in all models is the logarithm of mean real earnings in a group defined by micro-region, age-education cell and year. All nominal amounts are calculated in the currency as of January 2002.[3] In order to minimize potential problems of heteroskedasticity, we exclude cells containing fewer than 25 workers.

Throughout the estimation we use data only on the male population. This is restrictive, but it at least concentrates on a group whose labor-force participation is relatively unresponsive to wages and thus buttresses our treating quantities as exogenous. With this restriction (and, of course, with the exclusion of information on capital stocks by area—see Grant and Hamermesh, 1981), we are implicitly assuming that production is separable in inputs of male workers of various types from other inputs.

IV. Specification of Demand Systems

Let W be the logarithm of wages and X be an independent variable or vector of independent variables. Let i denote a micro-region, t time (Census year), and c an age-education cell. The simplest model is:

(1) $W_{itc} = \beta_{0c} + \beta_1 X_{itc} + \upsilon_{ic} + \theta_{tc} + \varepsilon_{itc}$, $i = 1,\ldots,K$; $t = 1,\ldots,T$,

where υ_i is a vector of area fixed effects and θ_t is a vector of time fixed effects. β_1 estimates the impact on wage rates of idiosyncratic variations in relative endowments of labor classified by age and education. The formulation in (1) restricts each own-quantity effect to be independent of

[3]To correct for currency changes, wages in 1970 and 1980 were divided by 2,750,000,000,000; and in 1991, they were divided by 2,750,000, as suggested by Corseuil and Foguel (2002). This correction was followed by the use of deflators suggested by the same authors. Both the correction for currency changes and the deflation are done for convenience only. Taking logarithms of wages, using nominal or real wages, of course generates the same estimates of the crucial parameters.

variations in the relative sizes of the other eleven age-education groups. We estimate Equation (1) in a single regression, including the proportions of people in each of the age-education groups, eleven indicators for age-education groups, and three for Census years. The reference group is workers ages 15-24 with zero to four years of schooling observed in the 1970 Census.

Because it sets all cross-quantity effects to zero, Equation (1) is highly restrictive. An approach that allows for cross-quantity effects, and thus accords more closely with theory by explicitly allowing labor-labor substitution, is:

(2) $W_{itc} = \beta_0 + \beta_1 X_{itc} + \beta_2 X_{itc'} + \upsilon_i + \theta_t + \varepsilon_{itc}$, $i = 1,\ldots,K;\ t = 1,\ldots,T,$

where c' refers to the other age-education cells. This formulation—a complete system of inverse labor-demand equations—allows for substitution parameters that indicate how a change in the fraction of the work force in one cell alters the wage rates of workers in any other cell. Equation (2) contains ten terms in $X_{itc'}$. We estimate a pooled version including all cross-proportions of people for each of the twelve age-education groups.[4]

Some of the restrictive assumptions in (1) and (2) can be relaxed still further. First, we can allow the production parameters to vary over time. With this relaxation (1) becomes:

(1') $W_{itc} = \beta_0 + \beta_1 X_{itc} + \beta_3 \theta_t X_{itc} + \upsilon_i + \theta_t + \varepsilon_{itc}$, $i = 1,\ldots,K;\ t = 1,\ldots,T.$

In this specification the X variables are interacted with the time indicators, which allows testing whether the own-quantity elasticities of factor price are unchanged over time. As such, (1') nests the simpler specification (1). A model analogous to (2) allows all the substitution parameters to vary over time by adding both own-quantity interactions with the T-1 time indicators and interactions of the cross-quantity terms with those indicators. It can be written:

(2') $W_{itc} = \beta_0 + \beta_1 X_{itc} + \beta_2 X_{itc'} + \beta_3 \theta_t X_{itc} + \beta_4 \theta_t X_{itc'} + \upsilon_i + \theta_t + \varepsilon_{itc}$, $i = 1,\ldots,K;\ t = 1,\ldots,T.$

[4]One could estimate both (1) and (2) without the area and time fixed effects, but we do not impose those restrictions anywhere in this study.

Note that (2') nests each of (1), (2) and (1').[5]

Throughout this series of specifications we have moved from the simplest formulation, Equation (1), to increasingly general formulations. Even the most complex specification (2') is not as general as it could be were still better data available. We have no *a priori* reason to specify constant marginal cost, but it is implicit in the formulations here and is necessitated by the absence of any information on the scale of production in each area. Similarly, absent data on size of firms we have assumed homotheticity throughout, and that assumption may also be too restrictive.

V. Descriptive Statistics and Estimates

A. Descriptive Analysis

As discussed above, the age distribution of the population of Brazil has been changing rapidly. Figure 1, based on UN estimates and projections, shows the evolution of the child- and old-age dependency ratios from 1950 to 2050. The child-dependency ratio has fallen dramatically and is projected to decrease significantly in the next decades. The old-age dependency ratio, however, has been increasing recently and will increase even more in coming years. These patterns are related to the decline in the total fertility rate since the 1960s (Table 1). Since fertility declined so abruptly, the shares of younger age groups have also declined.

Differences in the timing and speed of the fertility transition led to substantial temporal differences in the age distribution across regions, states and municipalities. Figure 2 illustrates the percentage of young adults (25-34 years of age) with at least nine years of schooling in all 502 Brazilian micro-regions for 1970-2000 censuses. There is a clear increase over time in the proportion of young adults with higher educational attainment. At the same time, differences among micro-regions are pronounced and persistent. Higher proportions in this age-education

---

[5]In all four of these models one could expand the specifications still further by allowing for time-varying area fixed effects, thus controlling for changing area-specific effects. Re-estimates of (1) and (2) suggest that this alternative did not affect the demand parameters on which we focus.

group are observed in the Southeastern (SE), Southern (SO), and Center-Western areas (CW) compared to the Northern (NO) and Northeastern (NE) areas. On the other hand, Figure 3 shows that the percentage of adults 35-49 with zero to four years of schooling has been decreasing in all micro-regions. Areas in the Southeast and South of Brazil show a greater decrease in the proportion of men in low-educated groups than do those in the North and Northeast.

Figure 4 presents age distributions in four selected micro-regions in 1970 and 2000. (Data are shown only for these two years to allow a clearer picture of the changes.) The curves for the Northeastern micro-regions (in the states of Piauí and Ceará) indicate that the age distributions in 1970 and 2000 were similar, unlike the Southeastern (Rio de Janeiro) and the Southern (Rio Grande do Sul) micro-regions, where the proportion in the older age groups grew from 1970 to 2000. Figure 5 illustrates the distribution of the male population by education for 1970 and 2000 for the same micro-regions. In general, the percentages of people with higher levels of schooling grew over the years. Furthermore, the Northeastern micro-regions have lower levels of education than those in the South and Southeast.

The crucial point is not only that there were profound demographic changes in Brazil over this period in both the age structure and educational attainment, but also that these proceeded at different rates in different parts of the country. These differences allow us to identify the labor-demand parameters—and thus to expand the study of the labor-market effects of demographic change generally. Also, the persistent differences in levels suggest the need to use models that account for specific local factors through the use of fixed effects for micro-regions.

B. Estimating the Effects on Labor-Market Outcomes

Since there are 502 micro-regions, twelve age-education groups and four Censuses, the maximum number of possible observations in the regressions is 24,096. The requirement that there be at least 25 individuals in a cell results, however, in excluding some observations, so that we use 19,727 observations throughout. Table 2 presents estimates of Equation (1). The indicator variables for age-education groups show that within each age category earnings are higher for those people with more schooling. For instance, men ages 25-34 with zero to four years of schooling earn 1.52 times what men ages 15-24 with same education (the reference category) earn. Young adults (25-34) with at least nine years of schooling earn 6.05 times more. The estimates are thus consistent with what we know about age-earnings profiles and the impact of education on them.

Table 3 presents estimates of (1’), which allows the estimated own-quantity effects to vary over time. Interactions with year indicators show that the negative impacts of the changing distributions of workers across micro-regions have been decreasing in some groups. This is observed mainly for 1991 and 2000 for the least educated and the oldest groups of workers. Among them the positive coefficients on the interaction terms essentially offset the negative coefficients on the main effect terms. Among the other groups, particularly the highest-educated and prime-age workers, the impacts of increasing shares of the work force remain negative throughout.

The estimated coefficients on the proportions of men in each age-education group (the β_1) generally indicate greater negative impacts as the amount of education embodied in a worker increases. In order to interpret these coefficients it is necessary to calculate elasticities, because the proportions of workers vary across age-education groups over time. In Table 4 we present the estimated own-quantity elasticities based on the parameter estimates in Table 3. The estimated elasticities of factor price generally tend to be more negative among age-education groups with more education (five to eight years of schooling, or at least nine years of schooling). Moreover, the negative impacts increase over time among the more educated groups, while the elasticities of factor price for groups with the lowest educational attainment become less negative over time. This latter result is especially interesting and might be explained by the increasing openness of the Brazilian economy and its exposure to competition from manufactured goods produced using even less skilled and less expensive labor than that of the least-educated Brazilians.

As can be seen in Table 4, an increase of ten percent in the number of people with five to eight years of schooling and between 15 and 24 years of age reduced their earnings by 2.7 percent in 1970 and 2.5 percent in 2000. Among young adults (25-34) in the same education group the impact of the same shock increases over time, from a reduction of 1.4 percent in 1970 to 2.8 percent in 2000. The same happens for experienced adults (35-49) and older adults (50-64) with five to eight years of schooling. Older adults (50-64) with at least nine years of schooling, however, and even experienced adults (35-49) in this education group, see smaller negative impacts on earnings.

Overall the results indicate that identical percentage-point increases in the proportion of people with more schooling generate larger negative impacts on earnings. This finding is consistent with substantial evidence from industrialized economies that own-price elasticities of labor demand fall with educational attainment and with the sparser evidence for those economies that elasticities of factor price rise with education (Hamermesh 1993, Chapter 3). Moreover, that the factor-price elasticities have generally been increasing over time except among the least-educated group suggests either that a given amount of education implies that more skill is embodied in a group of workers now than in 1970, or that the Brazilian labor market has become more rigid.

The estimates permit comparing the predicted mean monthly real earnings among different age-education groups over the range of the actual proportions of people in these groups in Brazilian micro-regions. Figure 6 shows predicted earnings for young men (25-34) with at least nine years of schooling, and for experienced men (35-49) with zero to four years of schooling. Most important, these figures illustrate that earnings are lower among men who live in micro-regions with higher proportions of male workers in their own age-education group. The negative impact on earnings of a higher proportion of workers in one's own group is greater for young adults with 9+ years of education, as can be seen by the steeper curves for them, than among adults with less education.

Another way to view the implications of the estimates of the own-effects model is to consider how the changing national distribution of males by age-education group from 1970 to 2000 affects predicted earnings in the group (Figure 7). To do this we use national proportions of males by age-education group and Census year to calculate predicted earnings, applying the parameter estimates from Table 3. The figure shows that groups with declining proportions experience gains in earnings, and vice-versa for groups whose representation is growing.

The ratio of predicted earnings from the own-effects model to predicted earnings from a model that includes only indicators for age-education group is plotted in Figure 7 for all three education groups for adult males ages 35-49. The horizontal line shows baseline predicted values from a model that only includes age-education indicators, i.e., that excludes the proportions in each age-education group and for comparison purposes indexes these at one. Comparing the curves in Figure 7 to the baseline predictions, one can see that the low-education group (zero to four years of schooling) has predicted earnings from the own-effects models that are increasing over time. These changes make it clear that we can predict earnings changes more accurately by taking into account the effects of changes in the relative size of each group on its earnings.

For each of the four selected micro-regions (those examined in Figures 4 and 5), in Figure 8 we show the ratio of predicted mean earnings from the simple model (1) that restricted cross-quantity effects to be zero to predicted earnings from the baseline regression that only includes the age-education group and year indicators. This Figure is designed to show the differences in estimated earnings between a model that takes into account age and education structures in a local area (proportions of people by age-education groups) and a model of that considers only the direct impact of age and education on earnings. A second set of ratios in Figure 8 compares predicted earnings from the model that includes cross-effects (that estimates (2)) and the baseline model that only includes age and education indicators.[6] The dashed line indicates the ratio between observed and predicted earnings from the model that only includes age-education indicators; it thus shows how well the predictions fit the data in the selected areas.

Figure 8 demonstrates that the slopes of predicted earnings from the own-effects and cross-effects models accord fairly well with the slopes actually found in the data, all in relation to predicted earnings based on a restrictive model that excludes group-size effects. Note, however, that the curves are flatter in the two Northeastern areas than in the two areas from the South and Southeast, due to the greater shift in the proportions in this group that occurred in those areas. Most important, the calculations presented in Figure 8 show that the variations observed over time generate patterns that are more similar to those in predicted earnings from the cross-effects model than from the more restrictive model that excludes them. This finding is clearest for the micro-regions in the Southeastern and Southern regions; but additional calculations for all regions suggest that this result is general, and thus that cross-effects are important.

We do not present estimates of predictions from the more complex model (2'). The general conclusion from those estimates, however, is that its additional complications do not alter the inferences we have made above: The estimated own-quantity elasticities are negative, and in general they are more negative the greater the skill embodied in a group of workers.

C. Robustness Considerations

Remembering that our purpose has been to examine the impacts of changing endowments of labor by type (in our case, by age and education), whether the strong evidence we have uncovered that there are substantial negative own-quantity elasticities of factor prices is robust to a number of potential problems is crucial. Consider perhaps the most important issue, inter-micro-regional migration. Ideally one might embed the production models estimated here in a more general model in which population is determined endogenously with wages through

[6]We do not present the estimates of the coefficients in (2) in tabular format because the number of coefficients (147 in total) is daunting and makes it hard to interpret the results.

endogenous flows of migrants and endogenous fertility. The difficulty with such an expanded model is that there is nothing in the underlying data that might identify any of the relationships; nor is there any instrument that one might use to remove endogeneity problems in the age-education distributions that we use here as regressors.

Absent any satisfactory instruments, consider the direction of the bias in the estimated own-quantity elasticities that might be induced by endogenous inter-area migration. Such migration presumably flows to those areas where relative declines in the size of the labor force in a particular age-education category have raised wage rates. If so, migration flows reduce wage rates in precisely those areas and among those age-education groups where natural (pre-migration) scarcities would have raised wages. Endogenous migration thus biases the estimated negative own-quantity effects on wages toward zero and implies, assuming migrants respond to relative wage differentials, that our estimates understate the absolute values of the effects.

Whether the sizes of the positive biases induced by endogenous migration differ by age-education category is a more difficult question and depends largely on differences in the responsiveness of migration to wage differentials by age-education category. Certainly we know that in Brazil, as elsewhere, migration is disproportionately among younger workers; and there is also evidence that Brazilian migration rates are highest among the least-educated. We do not know much, however, about differences in the responsiveness of migration to wages by age-education group. If, however, the elasticities of migration with respect to wages are, like labor-supply elasticities, greater for less-skilled workers, that would suggest that the positive biases in the own-quantity factor-price elasticities are greater for less-skilled workers—in this study, the young and the least-educated.

A similar problem might arise if educational attainment is endogenous—if within micro-regions young people attain more schooling when the returns to education have increased because more educated labor has become relatively scarce. If this occurs, the response would raise quantity endowments and reduce relative wages among those labor-force groups whose relative wage rates would otherwise have risen still further. Like the biases in own-quantity factor-price elasticities that may be induced by endogenous migration, those potentially induced by endogenous educational choices also reduce the absolute values of the (negative) estimated parameters below their true values.

Other potential difficulties might arise from our implicit assumption that the sub-aggregates of male labor are separable in production from those of other inputs, including capital and female labor. In the broad labor-demand literature formal tests of separability almost always reject that assumption (Hamermesh 1993, Ch. 3). Without data on capital stocks by area we cannot re-specify the model to account for this non-separability problem. The literature suggests quite clearly, however, that capital and skill are p-complements, so that we expect the returns to skill to be greater in those areas where production is more physical-capital intensive. Coupled with a probable positive correlation between skill intensity and endowments of physical capital, this correlation means that any mis-specification resulting from the absence of data on physical capital will bias the estimated own-quantity factor-price elasticities among more skilled workers toward zero.

We have the same Census information by micro-region on the age-education structure of the female workforce as we do for males. We could thus include these distributions as additional $X_{c'}$ variables in estimating (2), thus allowing for the cross-effects of relative quantities of female workers on male relative wages. The difficulties with doing this are several. Most important, the distributions of female workers are highly correlated with those of male workers. Where and when the male workforce is older, so is the female workforce; where and when men are better educated, so are women. How excluding quantities of female workers from the equations biases the estimated own-quantity factor-price elasticities for male workers depends on both this positive correlation and the substitution/complementarity relationships between male and female workers within and across age-education categories. There is no evidence on these latter correlations, so that in the end the direction of any bias is an empirical issue. We did, however, add the relative endowments of female workers to a set of re-estimates of (2); not surprisingly, the high positive correlations between the male and female distributions across the age-education categories within areas caused the standard errors to increase greatly, although the estimated elasticities did not change greatly.

Finally, while we are basing the estimates on local-area rather than establishment data, one might argue that we can proxy establishment size by area size and get a handle on possible problems in our results that occur if underlying production functions are heterothetic. To examine this we re-estimated all the models interacting area size with the proportion of workers in each age-education category (and including area size as a separate regressor). None of the vectors of interaction terms was statistically significant; the main-effect terms in the proportions of workers in the age-education cells remain statistically significant; and the inclusion of the interactions did not alter our inferences about the sizes or signs of the factor-price elasticities. This is an admittedly weak test, but it is the best possible with these data.

VI. Conclusions and Implications

In this study we have tackled an old question, but in a different context and with a different way of extracting lessons from the data. Interesting and important results concerning the effects of shifts in the age distribution of the working age population have been obtained by a series of authors who looked at this question in the context of the baby-boom generation's impacts on the earnings of different cohorts in the U.S. But the question has received little attention in the countries of Asia and Latin America, which are now experiencing substantial shifts in their age distributions due to large and rapid declines in fertility. In these countries, the shifts in the age distribution have also been accompanied by dramatic increases in educational attainment that one might expect would also alter earnings distributions. A major difference between the U.S. and Latin American countries such as Brazil is the magnitude of regional differences in the timing of both the educational and demographic transitions. These changes were fairly homogeneous across American states but have varied enormously geographically in

Brazil. This heterogeneity both motivates and enhances the value of our microeconomic geographical approach to the problem.

Our first and most important result is that relative group size matters. The own-quantity wage effects are generally negative, as predicted by factor-demand theory; and potential biases induced by a number of effects for which we could not adjust mean that, if anything, the true impacts of changing relative quantities are larger in absolute value than our estimates suggest. The results imply that workers classified by age-education group are not perfect substitutes, so that own cohort-education size generally depresses earnings. That the effects increase with education is consistent with the observation of lower own-wage elasticities as education increases. Also, while there may have been shifts in relative demand over the thirty-year time span that we have examined, unlike in the United States they do not appear to have been large enough to offset the effects of variations in relative supply.

The effects of biased technological change and/or institutional changes are suggested by the evidence generated by our models that allowed the technological parameters to vary over time. Some of the parameters become less negative, suggesting that more recent changes in relative supply have altered the relative wages of the least-skilled workers less than would have been the case in the 1970s. Indeed, the own-quantity effect among workers with 0-4 years of schooling is essentially zero, suggesting that the increasing relative scarcity of such workers is hardly contributing to an increase in the relative earnings of the (fewer) remaining workers in the group. The results also suggest that throughout the period the sharp increases in the relative supply of the most skilled workers have reduced relative wages in this group: Accounting for relative shifts in supply implies that wage inequality may have risen less than if these substitution effects had not occurred.

The main focus in the literature has been on the impacts of changes in the dependency ratio resulting from demographic change. In countries such as Brazil, this ratio is undergoing dramatic change and will continue to do so for several decades. It is also true, however, that in terms of both age and educational attainment the composition of the Brazilian labor force, and labor forces in developing countries generally, is undergoing dramatic shifts. Here we have investigated whether these compositional shifts have had effects beyond those that are normally analyzed in the literature using Mincerian earnings equations, and whether studying their role in the context of a formal theory of labor demand is worthwhile. Our results suggest that shifts in the demographic and skill structure of the labor force are indeed influential and that this approach represents a fruitful way of expanding the study of earnings and income inequality, a central problem in economic development.

References

Autor, D., L. Katz, and A. Krueger. 1998. "Computing Inequality: Have Computers Changed the Labor Market?" *Quarterly Journal of Economics*, 113: 1169-1214.

Berger, M. 1985. "The Effect of Cohort Size on Earnings Growth: A Reexamination of the Evidence." *Journal of Political Economy* 93: 561-573.

Bloom, D., and R. Freeman. 1986. "The Effects of Rapid Population Growth on Labor Supply and Employment in Developing Countries." *Population and Development Review*, 12: 381-414.

Bloom, D., R. Freeman and S. Korenman. 1987. "The Labour-Market Consequences of Generational Crowding." *European Journal of Population*, 3: 131-176.

Bloom, D., D. Canning, and J. Sevilla. 2003. *The Demographic Dividend: A New Perspective on the Economic Consequences of Population Change*. Santa Monica, CA: RAND.

Borjas, G. 2003. "The Labor Demand Curve is Downward Sloping: Reexamining the Impact of Immigration on the Labor Market." *Quarterly Journal of Economics*, 118: 1335-1374.

Borjas, G., R. Freeman, and L. Katz. 1997. "How Do Immigration and Trade Affect Labor Market Outcomes?" *Brookings Papers on Economic Activity*, 1997: 1-90.

Corseuil, C., and M. Foguel. 2002. "Uma Sugestão de Deflatores para Rendas Obtidas a Partir de Algumas Pesquisas Domiciliares do IBGE." Texto para Discussão, 897. Rio de Janeiro, RJ, Brazil: Instituto de Pesquisa Econômica Aplicada (IPEA).

Freeman, R. 1979. "The Effect of Demographic Factors on Age-Earnings Profiles." *Journal of Human Resources,* 14: 289-318.

Gindling, T., and D. Robbins. 2001. "Patterns and Sources of Changing Wage Inequality in Chile and Costa Rica During Structural Adjustment." *World Development*, 29: 725-7245.

Grant, J., and D. Hamermesh. 1981. "Labor Market Competition among Youths, White Women and Others," *Review of Economics and Statistics*, 63: 354-360.

Hamermesh, D. 1993. *Labor Demand*. Princeton, NJ: Princeton University Press.

Katz, L., and K. Murphy. 1992. "Changes in Relative Wages, 1963-1987: Supply and Demand Factors," *Quarterly Journal of Economics*, 107:35-78.

Mason, A. 2005. "Demographic Transition and Demographic Dividends in Developed and Developing Countries." Population Division, Department of Economic and Social Affairs, United Nations. UN/POP/PD/2005/2.

Potter, J.., C. Schmertmann, and S. Cavenaghi. 2002. "Fertility and Development: Evidence from Brazil." *Demography*, 39: 739-761.

Riani, J.L.R. 2005. "Determinantes do Resultado Educacional no Brasil: Família, Perfil Escolar dos Municípios e Dividendo Demográfico numa Abordagem Hierárquica e Espacial." Centro de Desenvolvimento e Planejamento Regional (CEDEPLAR/UFMG). Ph.D. Dissertation. Belo Horizonte, Brazil.

Roberts, M., and E. Skoufias. 1997. "The Long-Run Demand for Skilled and Unskilled Labor in Colombian Manufacturing Plants." *Review of Economics and Statistics*, 79: 330-334.

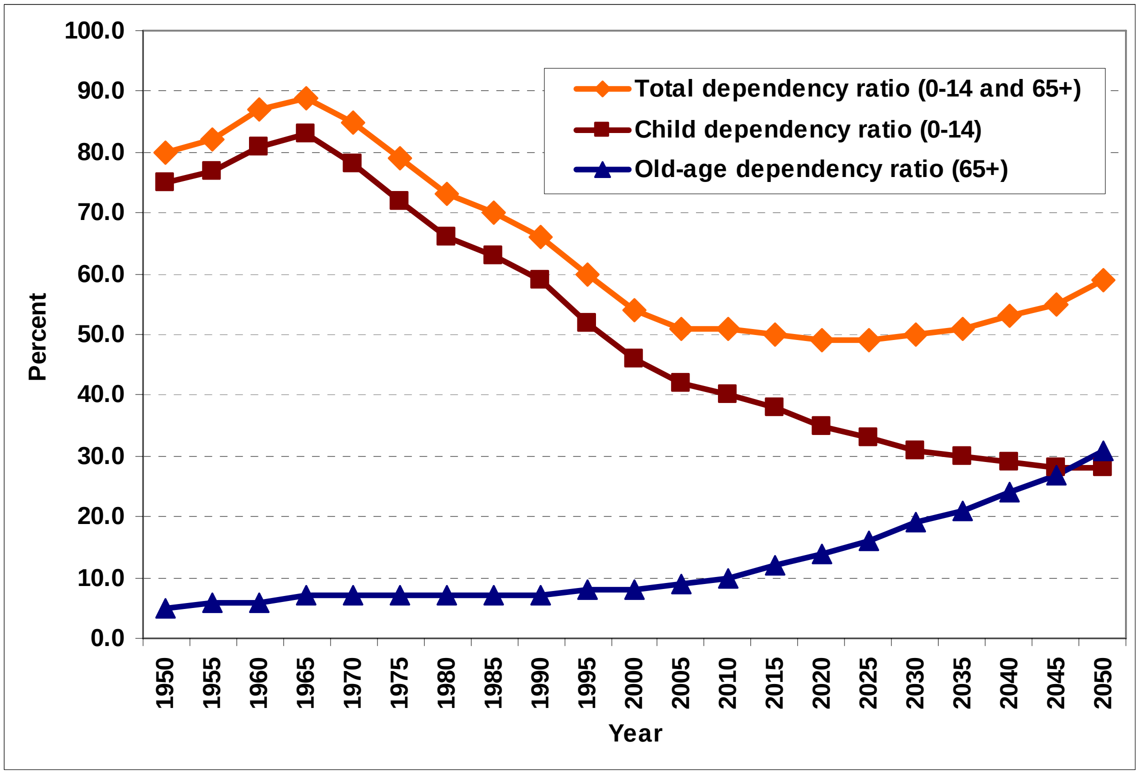
Saavedra, J., and M. Torero. 2004. "Labor Market Reforms and Their Impact over Formal Labor Demand and Job Market Turnover: The Case of Peru." In: *Law and Employment: Lessons from Latin America and the Caribbean.* (J. Heckman and C. Pagés, eds.). Chicago: University of Chicago Press, 131-182.

Triest, R., M. Sapozhnikov, and S. Sass. 2006. "Population Aging and the Structure of Wages." Chestnut Hill, MA: Boston College, Center for Retirement Research.

Welch, F. 1979. "Effects of Cohort Size on Earnings: The Baby Boom Babies' Financial Bust." *Journal of Political Economy*, 87: S65-S97.

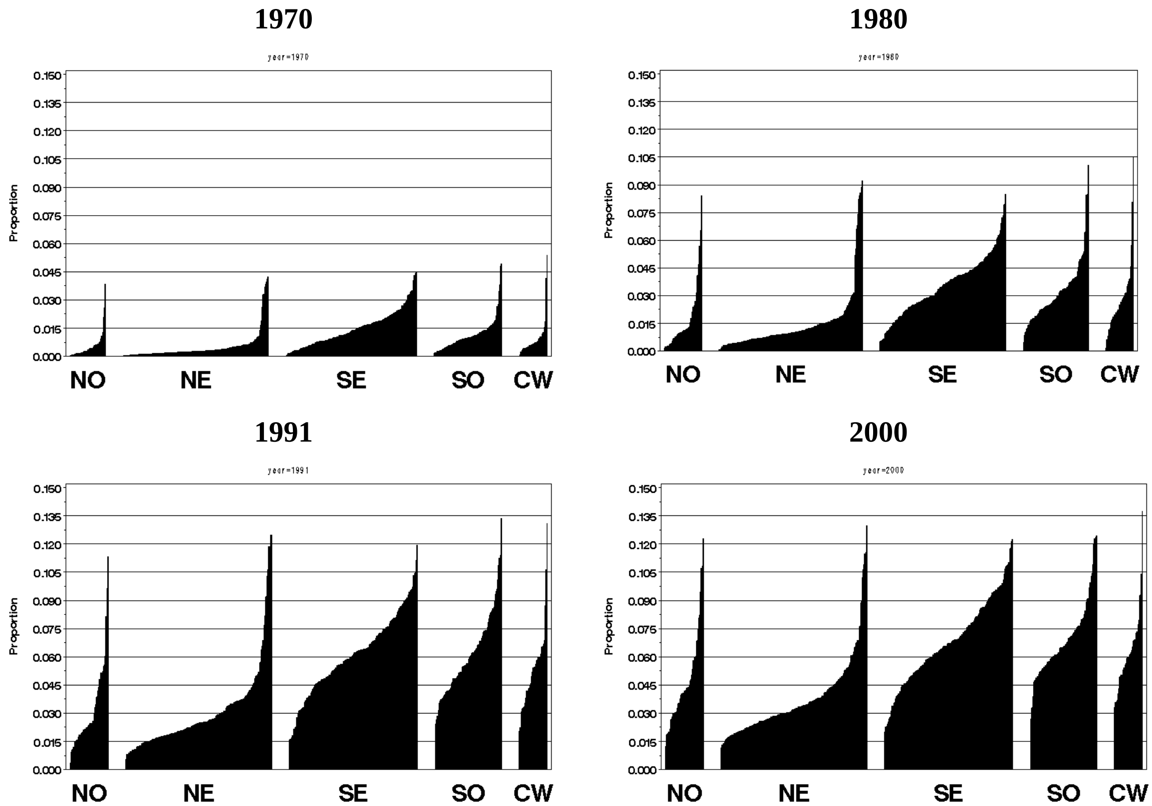
Williamson, J. 2003. "Demographic Change, Economic Growth, and Inequality." In: *Population Matters* (N. Birdsall, A.C. Kelley, and S.W. Sinding, eds.). Oxford: Oxford University Press, 106-136.

Figure 1. Dependency Ratios in Brazil, 1950–2050.



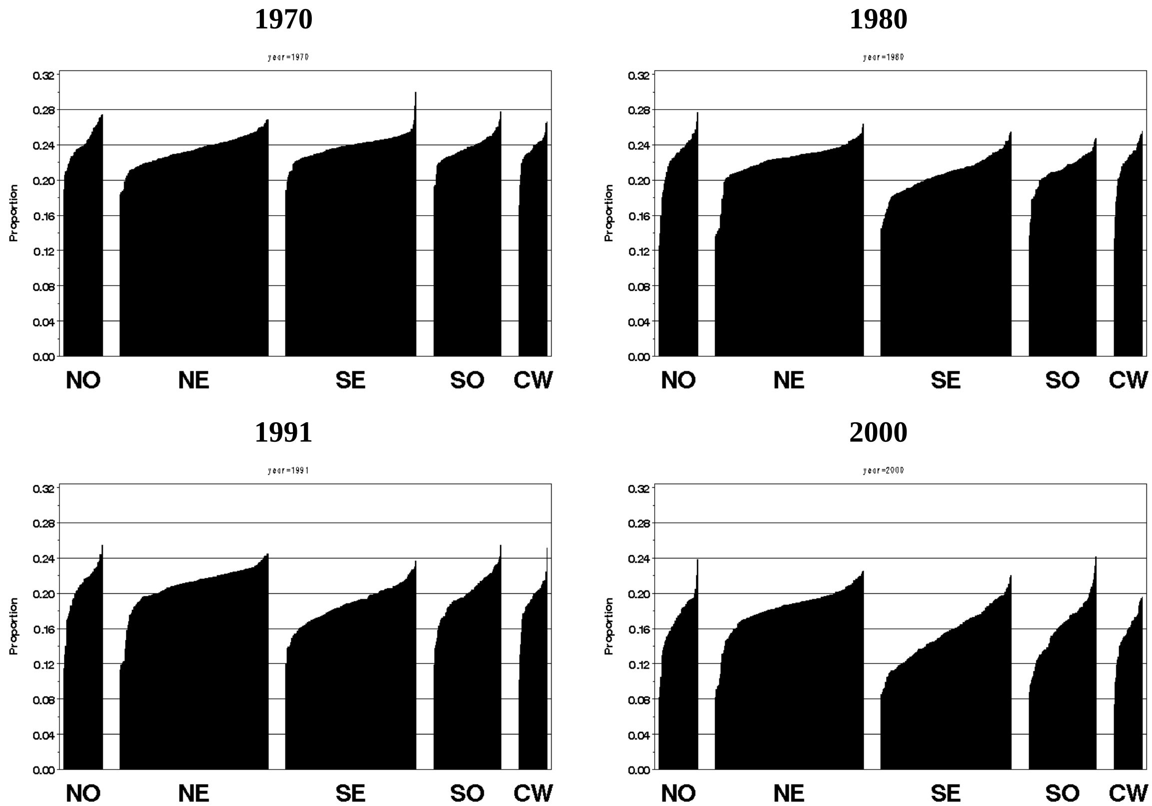
Source: United Nations - http://esa.un.org/unpp (in August 16, 2006 - medium variant).

Figure 2. Proportion of Men Ages 25–34 with 9+ Years of Schooling in 502 Micro-regions, 1970–2000.



Source: 1970–2000 Brazilian Censuses.

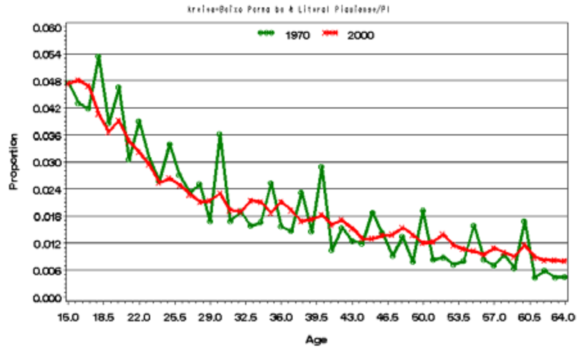
Figure 3. Proportion of Men Ages 35–49 with 0–4 Years of Schooling in 502 Micro-regions, 1970–2000.



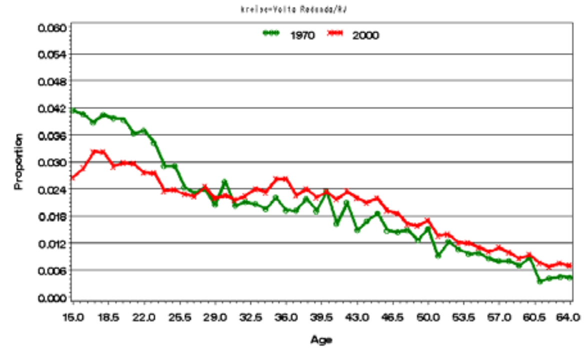
Source: 1970–2000 Brazilian Censuses.

Figure 4. Changes in the Age Distribution in Selected Micro-regions, 1970 and 2000.

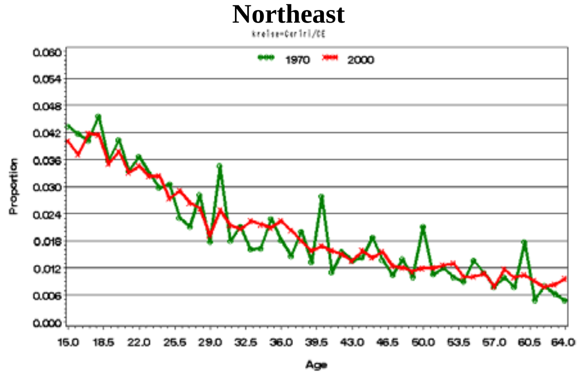
Baixo Parnaíba & Litoral Piauiense - Piauí
Northeast



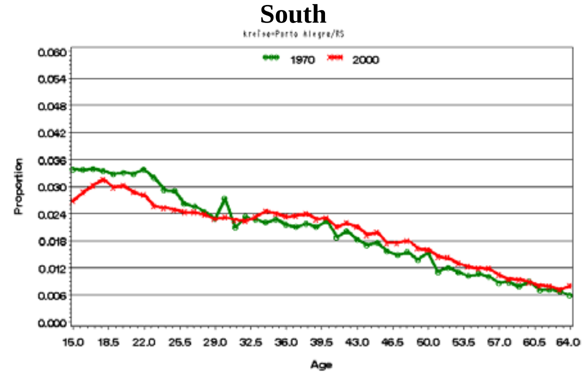
Volta Redonda - Rio de Janeiro
Southeast



Cariri - Ceará
Northeast



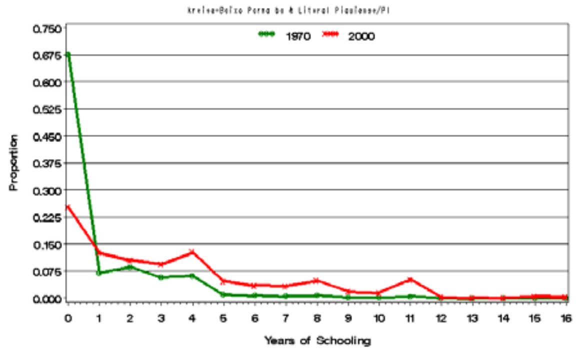
Porto Alegre - Rio Grande do Sul
South



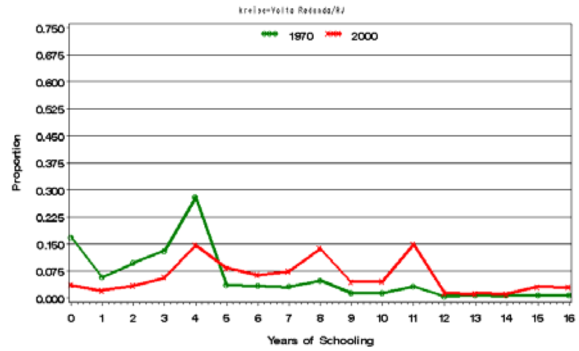
Source: 1970 and 2000 Brazilian Censuses.

Figure 5. Changes in the Education Distribution in Selected Micro-regions, 1970 and 2000.

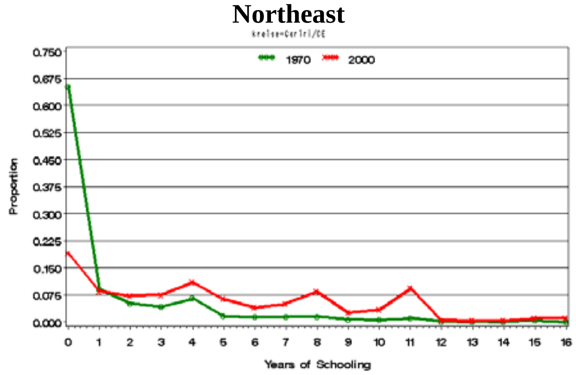
Baixo Parnaíba & Litoral Piauiense - Piauí
Northeast



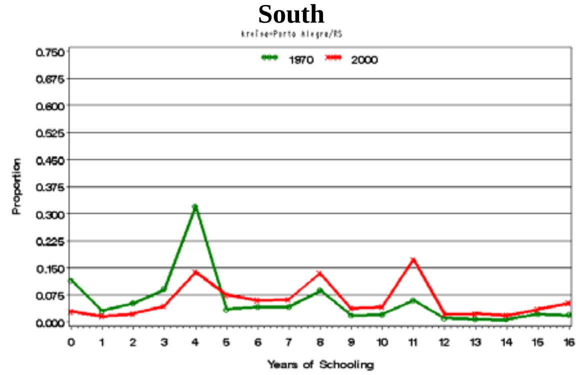
Volta Redonda - Rio de Janeiro
Southeast



Cariri - Ceará
Northeast

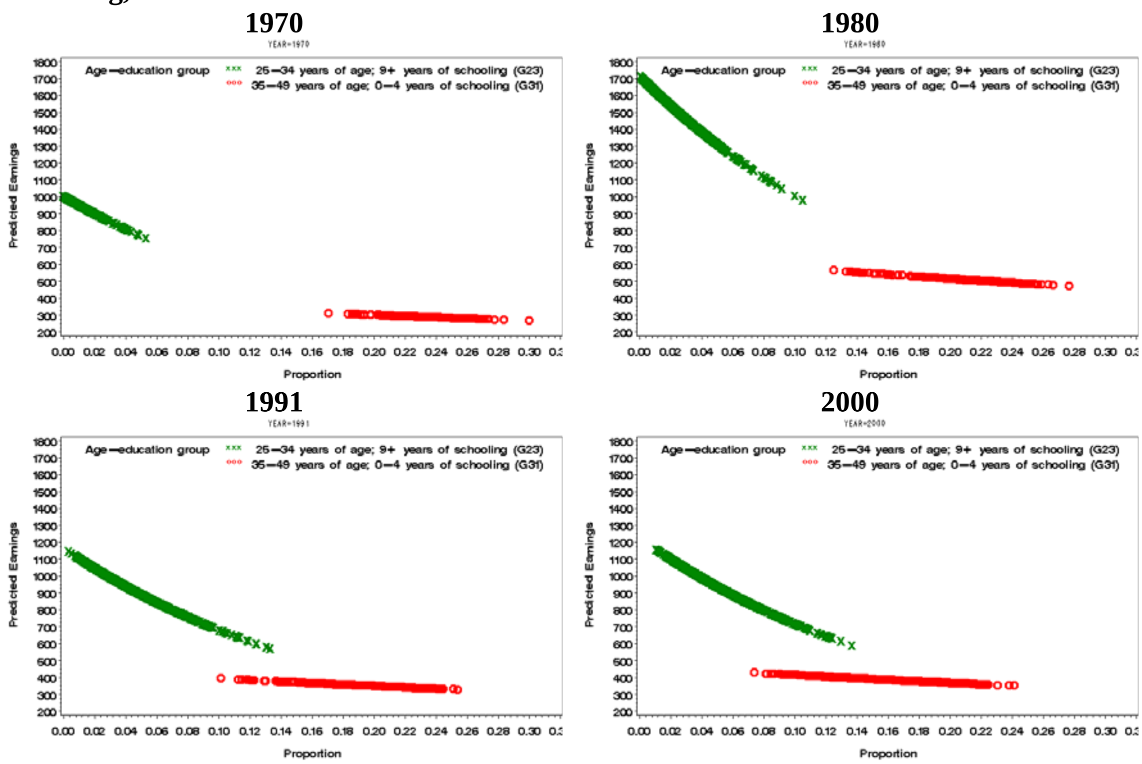


Porto Alegre - Rio Grande do Sul
South



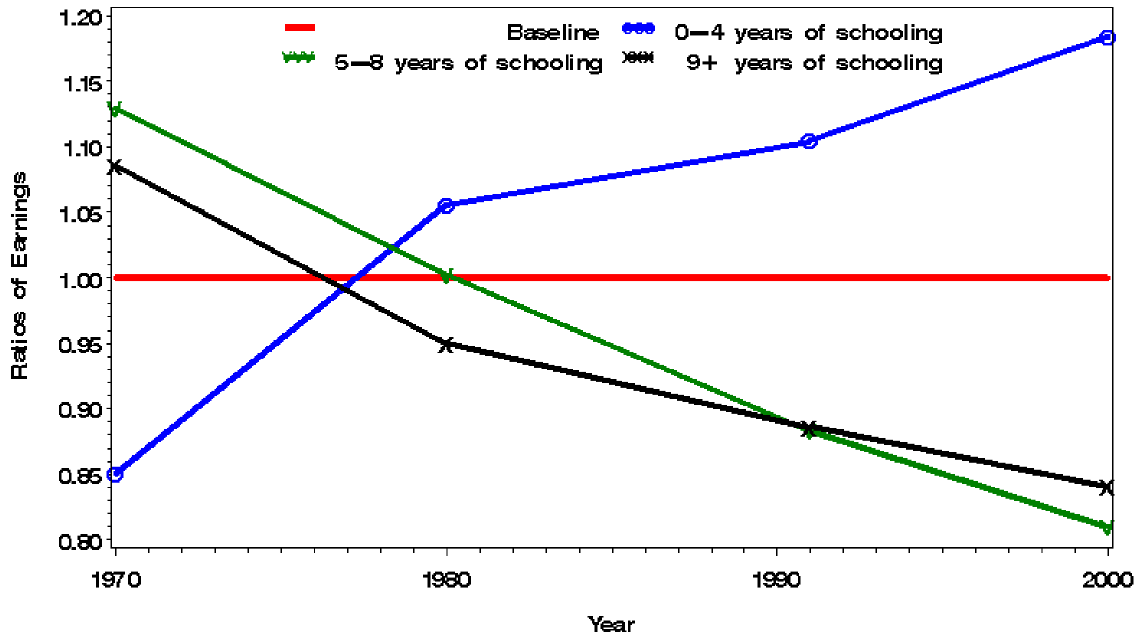
Source: 1970 and 2000 Brazilian Censuses.

Figure 6. Predicted Real Mean Monthly Earnings by Proportion of Population in Micro-regions, Men 25–34 with 9+ Years of Schooling, and Men 35–49 with 0–4 Years of Schooling, 1970–2000.



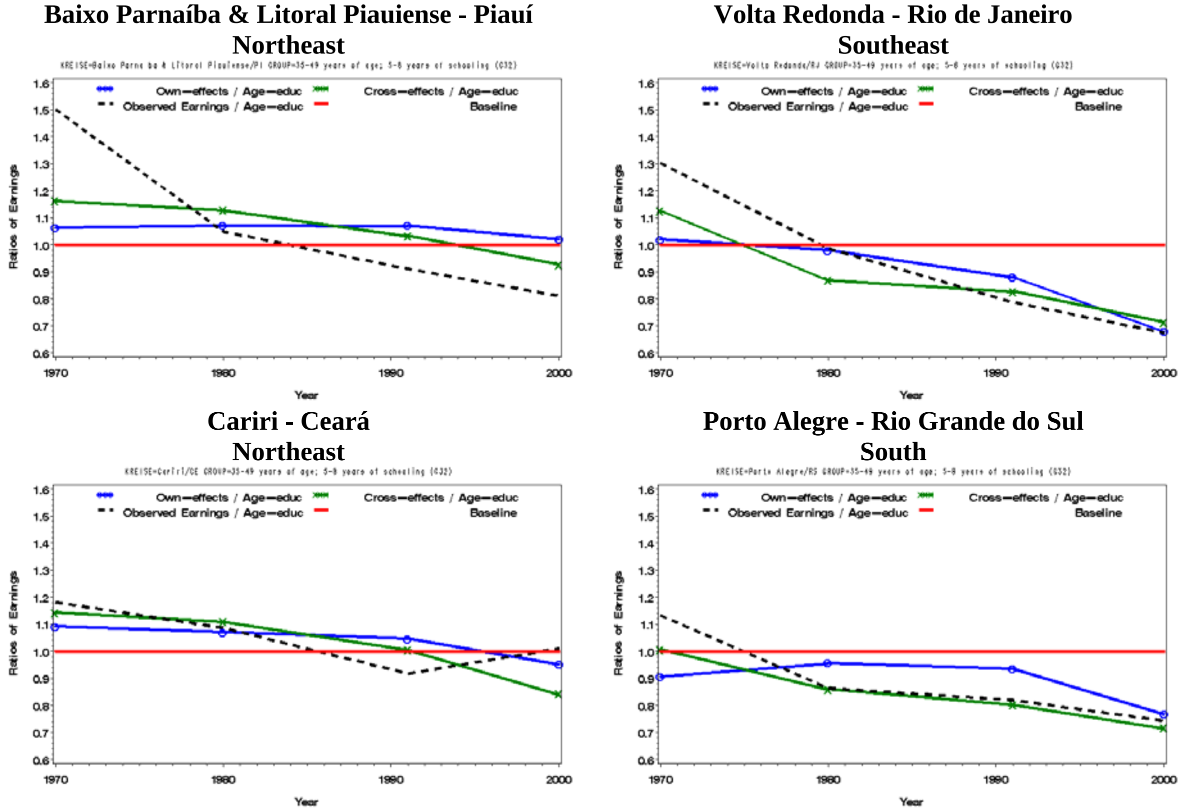
Source: 1970–2000 Brazilian Censuses.

Figure 7. Ratio of Predicted Mean Earnings from (1') to Predicted Earnings from Baseline Model, using the National Education Distribution for Males Ages 35–49, 1970–2000.



Source: 1970–2000 Brazilian Censuses.

Figure 8. Ratios of Observed Earnings, of Predicted Mean Earnings from (1), and of Predicted Mean Earnings from (2), to Predicted Earnings from Baseline Model, Males Ages 35–49 with 5–8 Years of Education, 1970–2000, Selected Micro-regions.



Source: 1970–2000 Brazilian Censuses.

Table 1. Total Fertility Rate, Infant Mortality Rate and Life Expectancy at Birth in Brazil, 1960–2005

Period	Total Fertility Rate	Infant Mortality (per 1,000 births)	Life Expectancy at Birth (years)
1960–1965	6.15	109.4	55.7
1965–1970	5.38	100.1	57.6
1970–1975	4.72	90.5	59.5
1975–1980	4.31	78.8	61.5
1980–1985	3.80	63.3	63.1
1985–1990	3.10	52.4	64.9
1990–1995	2.60	42.5	66.6
1995–2000	2.45	34.1	68.8
2000–2005	2.35	27.4	70.3

Source: United Nations - http://esa.un.org/unpp (in August 16, 2006).

Table 2. Fixed-Effects Estimates of (1), 1970–2000 (Dependent Variable is log(Monthly Earnings))

Variables	Coefficients	Elasticity (Based on the age-education distribution)			
		1970	1980	1991	2000
Constant	5.11***				
1970	—				
1980	0.54***				
1991	0.15***				
2000	0.20***				
Age-education Indicators					
15–24 years; 0–4 years of schooling	—				
15–24 years; 5–8 years of schooling	0.59***				
15–24 years; 9+ years of schooling	0.97***				
25–34 years; 0–4 years of schooling	0.42***				
25–34 years; 5–8 years of schooling	1.22***				
25–34 years; 9+ years of schooling	1.80***				
35–49 years; 0–4 years of schooling	0.83***				
35–49 years; 5–8 years of schooling	1.58***				
35–49 years; 9+ years of schooling	2.17***				
50–64 years; 0–4 years of schooling	0.83***				
50–64 years; 5–8 years of schooling	1.70***				
50–64 years; 9+ years of schooling	2.24***				
Proportions of Men in Age-education Groups					
Ages 15–24; 0–4 years of schooling	-0.08	-0.02	-0.02	-0.01	-0.01
Ages 15–24; 5–8 years of schooling	-3.32***	-0.18	-0.35	-0.40	-0.41
Ages 15–24; 9+ years of schooling	-4.81***	-0.13	-0.28	-0.29	-0.49
Ages 25–34; 0–4 years of schooling	-0.37**	-0.07	-0.06	-0.05	-0.03
Ages 25–34; 5–8 years of schooling	-6.00***	-0.12	-0.23	-0.41	-0.46
Ages 25–34; 9+ years of schooling	-5.37***	-0.11	-0.26	-0.40	-0.44
Ages 35–49; 0–4 years of schooling	-1.19***	-0.27	-0.23	-0.20	-0.16
Ages 35–49; 5–8 years of schooling	-7.23***	-0.12	-0.17	-0.27	-0.49
Ages 35–49; 9+ years of schooling	-3.08***	-0.05	-0.09	-0.17	-0.26
Ages 50–64; 0–4 years of schooling	-1.66***	-0.21	-0.20	-0.19	-0.17
Ages 50–64; 5–8 years of schooling	-16.12***	-0.11	-0.15	-0.19	-0.32
Ages 50–64; 9+ years of schooling	-0.25	0.00	0.00	0.00	-0.01
N observations	19,727				
N groups	502				
Fraction of variance due to the υ_i	0.73				
F (26; 19,199): All coefficients=0	8,506***				
F (501; 19,199): Area fixed effects=0	57.02***				

* Significant at p<.05; ** Significant at p<.01; *** Significant at p<.001, here and in Tables 3 and 5.

Table 3. Fixed-Effects Estimates of (1'), 1970–2000 (Dependent Variable is log(Monthly Earnings))

Variables	**Coefficients**			
Constant	5.30***			
1970	—			
1980	0.45***			
1991	-0.06***			
2000	-0.05***			
Age-education Indicators				
15–24 years; 0–4 years of schooling	—			
15–24 years; 5–8 years of schooling	0.52***			
15–24 years; 9+ years of schooling	0.90***			
25–34 years; 0–4 years of schooling	0.44***			
25–34 years; 5–8 years of schooling	1.12***			
25–34 years; 9+ years of schooling	1.68***			
35–49 years; 0–4 years of schooling	0.75***			
35–49 years; 5–8 years of schooling	1.51***			
35–49 years; 9+ years of schooling	2.11***			
50–64 years; 0–4 years of schooling	0.77***			
50–64 years; 5–8 years of schooling	1.61***			
50–64 years; 9+ years of schooling	2.23***			
		Interactions with Year		
Proportions of Men in Age-education Groups		**1980**	**1991**	**2000**
Ages 15–24; 0–4 years of schooling	-0.77***	0.34***	0.90***	1.32***
Ages 15–24; 5–8 years of schooling	-5.08***	0.71*	3.20***	3.06***
Ages 15–24; 9+ years of schooling	-4.82***	-1.18*	2.07***	1.68**
Ages 25–34; 0–4 years of schooling	-1.60***	0.98***	1.31***	1.63***
Ages 25–34; 5–8 years of schooling	-6.81***	0.20	2.87**	3.20***
Ages 25–34; 9+ years of schooling	-1.43	-2.39**	-0.93	-1.94*
Ages 35–49; 0–4 years of schooling	-1.98***	1.00***	1.56***	1.66***
Ages 35–49; 5–8 years of schooling	-8.67***	0.74	2.72*	3.92***
Ages 35–49; 9+ years of schooling	-4.63***	-1.04	3.67**	3.35**
Ages 50–64; 0–4 years of schooling	-3.44***	1.68***	2.67***	3.55***
Ages 50–64; 5–8 years of schooling	-8.66**	-2.38	-1.30	1.48
Ages 50–64; 9+ years of schooling	-15.98***	1.58	17.85***	19.39***
N observations	19,727			
N groups	502			
Fraction of variance due to the υ_i	0.74			
F (62; 19,163): All coefficients=0	3,957***			
F (501; 19,163): Area fixed effects=0	53.20***			

Table 4. Elasticities for Estimates of Equation (1'), using the National Age-Education Distribution, 1970–2000 (Dependent Variable is log(Monthly Earnings))

Proportions of Men in Age-education Groups	**Elasticities**			
	1970	**1980**	**1991**	**2000**
Ages 15–24 years; 0–4 years of schooling	-0.22	-0.09	0.02	0.05
Ages 15–24 years; 5–8 years of schooling	-0.27	-0.46	-0.23	-0.25
Ages 15–24 years; 9+ years of schooling	-0.13	-0.35	-0.16	-0.32
Ages 25–34 years; 0–4 years of schooling	-0.32	-0.10	-0.08	0.00
Ages 25–34 years; 5–8 years of schooling	-0.14	-0.26	-0.45	-0.28
Ages 25–34 years; 9+ years of schooling	-0.03	-0.18	-0.28	-0.27
Ages 35–49 years; 0–4 years of schooling	-0.45	-0.19	-0.17	-0.04
Ages 35–49 years; 5–8 years of schooling	-0.14	-0.19	-0.29	-0.32
Ages 35–49 years; 9+ years of schooling	-0.07	-0.16	-0.31	-0.11
Ages 50–64 years; 0–4 years of schooling	-0.44	-0.21	-0.20	0.01
Ages 50–64 years; 5–8 years of schooling	-0.06	-0.10	-0.13	-0.14
Ages 50–64 years; 9+ years of schooling	-0.10	-0.15	-0.25	0.10

Source: 1970–2000 Brazilian Censuses.